

Measuring the Impact of Consumption Tax on the Cost-of-Living Index from Japanese Household Survey

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Abstract

Japanese households experienced both the introduction of price consumption tax and two tax hikes from 1989 to 2014. According to a Bank of Japan report, the consumption tax hike (from 5% to 8%) in April 2014 was expected to increase the consumer price index by about two percentage points. In this study, we measure the impact of the consumption tax hike on the cost-of-living index, using the panel cointegrated demand system. We find that two consumption tax hikes in both 1997 and 2014 effectively raised the cost-of-living index. That is, it seems that the timing of two tax hikes was appropriate for Japanese households.

Keywords

Consumption Tax Hike, Cost-of-Living Index, Panel Cointegration, Demand System

1. Introduction

Japanese households experienced both the introduction of price consumption tax and two tax increases from 1989 to 2014. In 2012, the Japanese Prime Minister Shinzo Abe announced the “consumption tax increase” as one of his policy objectives under “Abenomics”¹, and it was enforced with an increase from 5% to 8% in April 2014². This policy objective is assumed as one of the key objectives leading to the future recovery of the Japanese economy.

It is wellknown that a consumption tax is a tax that a burden does not center on any particular person; it is collectively borne by the whole nation including

¹Abenomics is a popular name given for a series of economic policies in Japan that the politician Shinzo Abe of the Liberal Democratic Party advocated in the second Abe Cabinet.

²It has been announced that the consumption tax hike to 10% is postponed to October 2019.

the elderly. Therefore, the consumption tax is considered appropriate for social security resources in the aging Japanese society. In Japan, the consumption tax of 3% was first introduced in 1989, and which was next raised to 5% in 1997 and it was further raised to 8% in 2014. The Bank of Japan (BOJ) computed the direct effects of the tax increase with respect to the consumer price index (CPI). It showed that the consumption tax hike (from 5% to 8%) in April 2014 would increase the CPI by about two percentage points. Of course, as with other countries, the Japanese CPI itself includes the consumption tax in prices. In addition, it is difficult to measure the accurate influence of tax directly in CPI. Therefore, our estimation results also include the impacts of price rises by tax increase. **Table 1** compares the consumption tax with foreign countries. It is clear that the Japanese consumption tax is low compared to the EU zone, but is normal in Asian region. In general, we expect the household expenditure and prices to react more sensitively to a tax increase because the consumption tax is low.

In this study, we measure how the introduction of tax increase will raise or lower the cost-of-living index. We believe the reasons for focusing on the movement of the cost-of-living index are as follows: the price trend in Japan has always been a topic of interest, and it is advocated as the main policy for eliminating deflation. As the consumption tax increase is considered the basic for future policy, it is important to measure and predict its degree of influence on household budget and prices. Estimating the cost-of-living index could reveal the argument for a cost-of-living index is for the efficacy of inflation targeting policy, based on demand analysis.

To capture the influence of these price changes by consumption tax, we measure the cost-of-living index according to price changes and household budgets, and their fluctuations, using panel data for 26 years from 1989 to 2014. In our analysis, we use semi-macro panel data from 47 prefectural capitals. The panel data includes both time-series and cross-sectional dimensions and can measure the cost-of-living index for each prefectural capital.

The constitution of this paper is organized as follows. In Section 2, we introduce the almost ideal demand system (AIDS) model developed by Deaton and Muellbauer [1] and the price effects derived from this model. In Section 3, we

Table 1. The consumption tax for foreign countries in 2015.

Area	Country	Consumption tax (%)	Area	Country	Consumption tax (%)
EU	Denmark	25	Asia	Korea	10
	France	20		Taiwan	5
	Germany	19		New Zealand	15
	Netherlands	21		Japan	8
	Sweden	25		China	17
	Norway	25		Singapore	7

Note: The consumption tax indicates the value in 2015 and the source of reference is showed by national tax agency in Japan.

explain the data sources used in this study, and in Section 4, we report the estimation results, including the expenditure and price elasticity calculations. In Section 5, we calculate the cost-of-living index in Japan, including the effects of consumption tax increase during 1989 to 2014. Finally, in Section 6, we conclude the paper.

2. Model

2.1. Estimation Model

For this analysis, we use the AIDS model by Deaton and Muellbauer [1]. We use this model in panel cointegration analysis for time-series and cross-sectional dimensions [2]. Then, we accept the construction of the panel cointegration relationship for Japanese household expenditure. First, we note the price independence generalized logarithmic (PIGLOG) cost function as defined by

$$\ln C(u, \mathbf{p}) = \alpha(\mathbf{p}) + u \beta(\mathbf{p}) \quad (1)$$

where $\alpha(\mathbf{p})$ and $\beta(\mathbf{p})$ are functions of prices as follows.

$$\alpha(\mathbf{p}) = a_0 + \sum_{i=1}^n a_i \ln p_i + \frac{1}{2} \sum_{i=1}^n \sum_{j=1}^n b_{ij}^* \ln p_i \ln p_j$$

$$\beta(\mathbf{p}) = b_0 \prod_i p_i^{b_i}.$$

The cost function in (1) is homogeneous in \mathbf{p} . Next, the i -th budget share can be derived from $\partial \ln C / \partial p = \mathbf{W}$ and expressed by

$$w_{ikt} = a_i + \sum_{j=1}^n b_{ij} \ln p_{jkt} + c_i \ln \left(\frac{x_{kt}}{P_{kt}} \right) + u_{ikt}, \quad (2)$$

$$i, j = 1, \dots, n, k = 1, \dots, K, t = 1, \dots, T,$$

where n is the number of commodities in the system, w_{ikt} denotes the i -th budget share at the individual k in period t , $\ln p_{jkt}$ is the log price of commodity i at the individual k in period t , and $\ln \left(\frac{x_{kt}}{P_{kt}} \right)$ is the log real expenditure with $\ln P_{kt}$ of the aggregate price index. Originally, $\ln P_{kt}$ is given by

$$\ln P_{kt} = a_0 + \sum_{i=1}^n a_i \ln p_{ikt} + \frac{1}{2} \sum_{j=1}^n \sum_{i=1}^n b_{ij} \ln p_{ikt} \ln p_{jkt}. \quad (3)$$

In this analysis, we use $\ln P_{kt}$ linearly approximated by Stone's [3] form³ in substitution for (2):

$$\ln P_{kt} = \sum_{i=1}^n w_{ikt} \ln p_{ikt} \quad (4)$$

That is, we estimate the linearly approximated model in this study.

Furthermore, in our panel data, the error term u_{ikt} in (2) can be written as

$$u_{ikt} = \theta_{ik} + \mu_{it} + e_{ikt}, \quad (5)$$

where θ_{ik} denotes an individual fixed effect and μ_{it} denotes a time fixed effect. Further, e_{ikt} is usually assumed to have strong exogeneity and

³The linearly approximated AIDS model has often been used in the cointegration analysis because using the nonlinear price index of (3) is to complicate of estimation problem for linear combinations [4] [5].

$$E(\mathbf{e}|\boldsymbol{\theta}, \boldsymbol{\mu}, \ln \mathbf{p}, \ln(\mathbf{x}/\mathbf{P})) = 0.$$

The AIDS model requires satisfying the adding-up, homogeneity, and symmetry conditions in the parameters. The adding-up condition, which is automatically satisfied by the use of the $n-1$ equations in the estimation, is

$$\sum_{i=1}^n a_i = 1 \quad \text{and} \quad \sum_{i=1}^n b_{ij} = \sum_{i=1}^n c_i = 0.$$

The homogeneity restriction for price parameters is

$$\sum_{j=1}^n b_{ij} = 0, \quad (6)$$

and the symmetry restriction is

$$b_{ij} = b_{ji}. \quad (7)$$

Both these restrictions are imposed on price parameters in the estimation.

The expenditure elasticity of commodity i in individual k with respect to log real expenditure is given by

$$\eta_{ik} = 1 + \frac{c_i}{w_{ik}}. \quad (8)$$

Marshallian price elasticity with respect to log prices⁴ is given by

$$\lambda_{ijk} = -\delta_{ij} + \frac{b_{ij} - c_i w_{jk}}{w_{ik}}, \quad (9)$$

where δ_{ij} represents the Kronecker delta and is 1 when $i = j$ and is 0 otherwise.

2.2. Cost-of-Living Index

We consider the evaluation of cost to a price change. The cost-of-living index is defined as the ratio of the minimum expenditure required to attain the base preference at prices \mathbf{p}^0 to that required at prices \mathbf{p}^1 . We measure the consumer surplus for a price change from \mathbf{p}^1 to \mathbf{p}^0 by the cost function of (1) as follows:

$$\ln C(u, \mathbf{p}^1) - \ln C(u, \mathbf{p}^0) = \ln \frac{\alpha(\mathbf{p}^1)}{\alpha(\mathbf{p}^0)} + u^0 \ln \frac{\beta(\mathbf{p}^1)}{\beta(\mathbf{p}^0)}, \quad (10)$$

where u^0 is the base utility level and equals $\ln[x^0/\alpha(\mathbf{p}^0)]/\ln[\beta(\mathbf{p}^0)/\alpha(\mathbf{p}^0)]$. This is called the cost-of-living index. Equation (10) is specific to the PIGLOG model and shows how the price index varies with the households' standard of living. The first term is expressed by the price index change for a price change from \mathbf{p}^1 to \mathbf{p}^0 . In addition, the second term is expressed by the base utility level and a price change from \mathbf{p}^1 to \mathbf{p}^0 . There is available literature for the cost-of-living index, such as that provided by Deaton and Muellbauer [1], Fry and Parshardes [8], Lewbel [9], and Pollak [10]. For example, Fry and Parshardes [8] constructed a true cost-of-living index of the PIGLOG model by modelling substitution as an aggregation shift parameter and

⁴We use the price elasticity formula that is suitable for linear approximation model. It has been pointed out that using the original calculation formula occurs large distortion price elasticity [6] [7].

exploiting the Tornqvist index.

3. Data and Source

In this analysis, we use semi-macro panel data because the nonstationary problem is obvious in a time-series dimension and price effects can be accurately estimated using panel data. The household survey data include the panel data for workers' households in 47 prefectural capitals. The source of data is the *Family Income and Expenditure Survey* ("Kakei Chosa" in Japanese) by the Japanese Statistics Bureau from 1989 to 2014. We classified the data into 10 goods: food, housing, fuel, furniture, clothing, medicine, transportation, education, recreation, and miscellaneous. Price series data are obtained from the CPI and are calculated using 2010 as the base year⁵. As described in section 1, we know that it is difficult to eliminate the influence of tax directly in CPI.

Figure 1 and **Figure 2** show the average budget shares and log prices over the period 1989 to 2014 for 47 prefectural capitals. In the budget shares in **Figure 1**, several commodities including fuel, medicine, transport, and education display an upward trend for 26 years. In particular, the growth rate for transport is the largest among these commodities, capturing the rapid increase of communication expenses on Internet and mobile phone usage included under the transport category in recent years. However, the budget shares for housing, clothing, and miscellaneous display a downward trend. These commodities are those that have been classified as luxury goods in Japan. Moreover, the budget shares for these commodities have decreased for the last 26 years due to economic fluctuations. That is, Japanese households have reduced the budget shares for luxury goods and have increased the budget shares for necessity goods. In **Figure 2**, we find a remarkable change in log prices. The log prices for housing and recreation display a downward trend. In particular, among the 10 commodities, the reduction for housing is the most remarkable, and this change would influence the decline

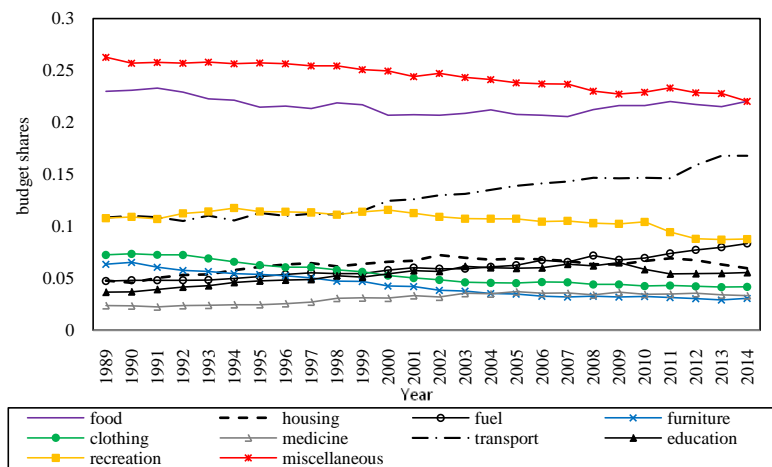


Figure 1. Budget shares from 1989-2014.

⁵The Japan's Statistics Bureau has revised the base year of CPI every five years. We used 2010 as the base year, with the closest price level to the present.

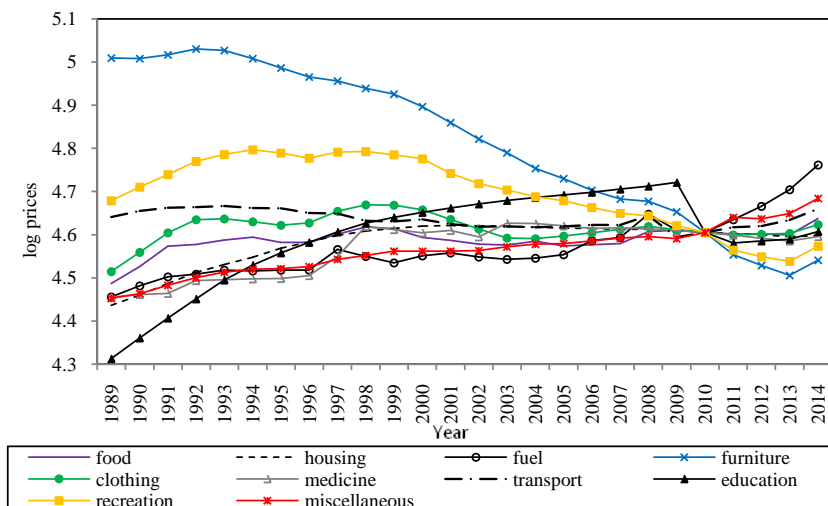


Figure 2. Log prices from 1989-2014.

of own-budget shares. In contrast, the log prices for fuel, medicine, education, and miscellaneous display an upward trend. In particular, the increase for education is noteworthy, and this could be related to improvements in the level of education in Japan over the past 20 years.

When we observe the long-term trend of these prices, the change that shows the influence of tax increase appears in 2014. In addition, when we specifically focus on the prices for fuel, clothing, medicine, and recreation, the price change due to the increased tax can be observed in 1997 after its introduction. As can be seen in (10), the price of movement is important when we consider the effect of the tax increase on the cost-of-living index.

4. Empirical Estimation

4.1. Panel Unit Root and Cointegration Tests

We know that budget shares lie between 0 and 1, and, therefore, cannot remain absolutely non stationary. Nevertheless, budget shares can closely approximate a nonstationary process. In fact, many previous studies on time-series have shown that budget shares have a nonstationary process, integrated with order one in unit root tests [11] [12]. In addition, we know that budget shares have a panel nonstationary process, including both time-series and cross-sectional dimensions [2].

First, **Table 2** reports the panel unit root test results for budget shares, log prices, log relative prices, and real expenditure in the Fisher ADF-type test by Maddala and Wu [13]. In most variables, we assume the individual effects and trend in the ADF regressions. The result shows that budget shares are integrated with order one, $I(1)$ as in Ogura [2]. In addition, most log prices, log relative prices, and log real expenditure are integrated with order one, $I(1)$. That is, we are able to reject the null hypothesis of no unit root in level at the 5% significance level. However, log prices for fuel and miscellaneous cannot reject the null

hypothesis of no unit root. In the framework of AIDS model, when the homogeneity restrictions are imposed in (2), log price terms are re-expressed by log relative prices. Therefore, it is also important that log relative prices have a panel unit root process. We find that all log relative prices reject the null hypothesis of no unit root.

Next, we confirm the existence of a long-term relationship between budget shares, log relative prices, and log real expenditure in panel data. We examine the panel cointegration test according to the economic relationship. Stationary residuals for equations would require these models to be cointegrated for each commodity i . **Table 3** reports the panel cointegration test results following Kao's [14] residual-based cointegration test. We confirm the cointegration relationships for residuals. We assume the null hypothesis of no cointegration relationship in the ADF regression, and find that the null hypothesis is rejected for all residuals at the 5% significance level. That is, we reveal the evidence of panel cointegration in the long-term relationship. If we reveal multiple cointegration relationships, it is important to show the type of cointegration relationship that exists between the nonstationary variables. In particular, we must indicate whether there is a linear combination of budget share, log relative prices, and log

Table 2. Panel unit root test.

Variables	Test statistics	P -value	Variables	Test statistics	P -value	Variables	Test statistics	P -value
w_1	211.482	0.0000	$\ln P_1$	295.928	0.0000	$\ln \frac{P_1}{P_{10}}$	288.696	0.0000
w_2	326.555	0.0000	$\ln P_2$	199.748	0.0000	$\ln \frac{P_2}{P_{10}}$	154.699	0.0001
w_3	238.643	0.0000	$\ln P_3$	24.0985	1.0000	$\ln \frac{P_3}{P_{10}}$	127.924	0.0115
w_4	354.030	0.0000	$\ln P_4$	158.952	0.0000	$\ln \frac{P_4}{P_{10}}$	142.172	0.0010
w_5	231.410	0.0000	$\ln P_5$	255.060	0.0000	$\ln \frac{P_5}{P_{10}}$	211.546	0.0000
w_6	297.395	0.0000	$\ln P_6$	121.658	0.0290	$\ln \frac{P_6}{P_{10}}$	148.9.3	0.0003
w_7	513.374	0.0000	$\ln P_7$	153.968	0.0001	$\ln \frac{P_7}{P_{10}}$	124.486	0.0193
w_8	281.999	0.0000	$\ln P_8$	465.842	0.0000	$\ln \frac{P_8}{P_{10}}$	143.920	0.0007
w_9	297.726	0.0000	$\ln P_9$	265.695	0.0000	$\ln \frac{P_9}{P_{10}}$	172.324	0.0000
w_{10}	369.518	0.0000	$\ln P_{10}$	110.717	0.1148	$\ln \frac{P_{10}}{P_{10}}$	288.696	0.0000
$\ln \frac{x}{P}$	385.942	0.0000						

Notes: The number of subscript in variables is corresponding to commodity: 1) food, 2) housing, 3) fuel, 4) furniture, 5) clothing, 6) medicine, 7) transport, 8) education, 9) recreation, and 10) miscellaneous. $\ln P_6$, $\ln P_7$, $\ln P_8$, $\ln P_9/P_{10}$, and $\ln P_{10}/P_{10}$ assume no trend in the ADF regressions. In addition, $\ln P_2/P_{10}$ and $\ln P_6/P_{10}$ assume no trend and no intercept in the ADF regressions. Other variables assume a trend and individual effects in the regressions.

Table 3. Panel cointegration test.

Residuals	Test statistics	P-value
u_1	-12.8094	0.0000
u_2	-6.6365	0.0000
u_3	-8.2214	0.0000
u_4	-16.7947	0.0000
u_5	-8.8651	0.0000
u_6	-11.4385	0.0000
u_7	-9.2008	0.0000
u_8	-8.3629	0.0000
u_9	-10.0427	0.0000

Notes: The number of subscript in each residual is corresponding to commodity: 1) food, 2) housing, 3) fuel, 4) furniture, 5) clothing, 6) medicine, 7) transport, 8) education, 9) recreation. We assume no deterministic trend in all regressions.

real expenditure that achieves stationarity. Even if we reveal long-term relationships in different sub-combination, it is not economically meaningful for our analysis. Therefore, we perform a residual-based cointegration test in each budget share equation.

4.2. Estimation of Panel Cointegration Relations in Demand Systems

When we estimate a single cointegration equation where there is at most one cointegration relationship among $I(1)$ variables, we may use the panel fully-modified estimator. However, we need to estimate the $r = n - 1$ cointegration relationships in a demand system framework. In addition, because the Slutsky symmetry in a demand system has cross-equation restrictions, we are also required to estimate the number of equations, simultaneously imposing these restrictions on cointegration vectors⁶. Therefore, we use the panel triangular error correction model (TECM) in our estimation, and apply the TECM suggested by Philips [15] [16] to the framework of a panel cointegrated demand system. This model can be expressed by

$$\mathbf{y}_{1t} = \mathbf{B}\mathbf{y}_{2t} + \boldsymbol{\varepsilon}_{1t}, \quad (11)$$

$$\Delta\mathbf{y}_{2t} = \boldsymbol{\varepsilon}_{2t}, \quad (12)$$

where \mathbf{y}_{1t} is $n_1 \times 1$ vector of left-hand side variables of the $n - 1$ cointegration system and \mathbf{y}_{2t} is $n_2 \times 1$ vector of right-hand side variables. In addition, $\boldsymbol{\varepsilon}_{1t}$ is $n_1 \times 1$ subvector, and $\boldsymbol{\varepsilon}_{2t}$ is $n_2 \times 1$ subvector. The cointegration parameters \mathbf{B} is the $n_1 \times n_2$ matrix, and then $\mathbf{y}_{1t} = \mathbf{B}\mathbf{y}_{2t}$ represents the linear long-run relationship of demand system, and $\boldsymbol{\varepsilon}_{1t}$ represents the short-run deviations from long-run equilibrium in (11), *i.e.* a measure of the difference between observed budget shares \hat{w}_{ikt} and theoretical shares w_{ikt} . Philips [15] [16] also showed that the maximum likelihood estimator of the cointegration

⁶When we impose only the homogeneity restrictions on the cointegration vector, it is possible to estimate in a single Equation.

parameters B is equivalent to the OLS estimator as the following augmented regression Equations⁷:

$$y_{1t} = B y_{2t} + \sum_{l=0}^L C_l \Delta y_{2t-l} + v_t. \tag{13}$$

In addition, Philips [15] [16] considered demand homogeneity tests for q restrictions on the cointegration parameters matrix B . When the null hypothesis is denoted as

$$H_0 : R_1 B r_2 = r,$$

where R_1 is $q \times n_1$ matrix, r_1 is $n_2 \times 1$ vector, and r_2 is $q \times 1$ vector. He also showed that the Wald test is valid for testing the hypothesis on the elements B . The form of Wald test statistics is given by

$$W = \frac{(R_1 B r_1 - r_2)' (R_1 \tilde{\Sigma} R_1')^{-1} (R_1 B r_1 - r_2)}{r_2' (X' M X)^{-1} r_2}, \tag{14}$$

where $\tilde{\Sigma} = n^{-1} \tilde{V} \tilde{V}'$. But when there is serial correlation in equation error, it can be replaced by the variance-covariance matrix of footnote 7.

Table 4 shows the Wald test results for the various restrictions, calculated by the estimated cointegration vectors in the panel TECM. In (13), we select the augmented lag $L=0$ by the Bayesian Information Criterion (BIC). First, we test the price effects, $H_0 : b_{ij} = 0$ (for $i, j = 1, \dots, 9$) and log real expenditure effects, $H_0 : c_j = 0$ (for $j = 1, \dots, 9$) in the long-run vectors. These null hypotheses are all rejected at the 5% significance level and we find that the price effect and the expenditure effect are statistically significant. Next, we consider that the homogeneity and the Slutsky symmetry tests including homogeneity restrictions. We use the null hypothesis in (6) and (7) and attempt the Wald test against the alternative. If the null hypothesis is rejected, the homogeneity and Slutsky symmetry restrictions on cointegration vectors are not established statistically. That is, the consistency between demand theory and the data in long-run relation-

Table 4. Model specification tests in cointegration relations.

Specifications	Null hypothesis	df	Test statistics	P-value
Price effect	$b_{ij} = 0$ (for $i, j = 1, \dots, 9$)	81	6639.539	0.000
Expenditure effect	$c_j = 0$ (for $j = 1, \dots, 9$)	9	885.459	0.000
	$b_{ij} = 0, c_j = 0$ (for $i, j = 1, \dots, 9$)	90	780.091	0.000
Homogeneity test	$\sum_j b_{ij} = 0$ (for $j = 1, \dots, 9$)	9	8.027	0.531
Symmetry test (including homogeneity)	$b_{ij} = b_{ji}$ (for $i, j = 1, \dots, 9$)	45	35.593	0.841

⁷Note that the t-value derived from the OLS estimator should be corrected by the Newey-West [17] estimator of the variance-covariance matrix because there may exist both heteroscedasticity and autocorrelation in equation errors. The variance-covariance matrix

$$\tilde{\Sigma} = \tilde{\Omega} + \sum_{s=1}^q f(s, q) [\tilde{\Omega}_s + \tilde{\Omega}'_s],$$

where $\tilde{\Omega}_s = T^{-1} \sum_{s=2}^T v_t v'_{t-s}$, and $f(s, q) = 1 + [s/(q+1)]$.

Table 5. Estimated long-run price and expenditure elasticities in demand.

Commodity	Marshallian price elasticity										Expenditure elasticity
	food	housing	fuel	furniture	clothing	medicine	transport	education	recreation	miscellaneous	
1. food	-0.0202 (0.0020)	-0.0676 (0.0003)	-0.0710 (0.0002)	-0.0578 (0.0002)	-0.0593 (0.0002)	0.0156 (0.0001)	-0.0588 (0.0003)	-0.0187 (0.0002)	0.0451 (0.0002)	-0.2563 (0.0006)	0.5490 (0.0010)
2. housing	-0.1498 (0.0013)	0.2800 (0.0112)	0.1262 (0.0012)	-0.0339 (0.0005)	0.0083 (0.0004)	-0.0899 (0.0009)	-0.1003 (0.0013)	-0.1327 (0.0014)	-0.0838 (0.0008)	0.1599 (0.0021)	0.0159 (0.0090)
3. fuel	-0.2341 (0.0013)	0.1015 (0.0006)	-0.7569 (0.0011)	-0.0901 (0.0004)	-0.2031 (0.0010)	-0.1479 (0.0009)	-0.3802 (0.0023)	-0.1507 (0.0009)	-0.3951 (0.0021)	-0.0288 (0.0005)	0.4080 (0.0032)
4. furniture	-0.3981 (0.0032)	-0.1048 (0.0008)	-0.1659 (0.0013)	-0.3533 (0.0052)	-0.3083 (0.0025)	-0.0323 (0.0003)	0.2627 (0.0022)	-0.0111 (0.0001)	-0.3096 (0.0025)	-0.1760 (0.0014)	0.9055 (0.0008)
5. clothing	-0.3720 (0.0024)	-0.0571 (0.0004)	-0.2752 (0.0018)	-0.2462 (0.0015)	0.1990 (0.0077)	0.0501 (0.0003)	-0.2222 (0.0015)	-0.0039 (0.0001)	0.1112 (0.0007)	-0.4194 (0.0027)	1.1012 (0.0006)
6. medicine	0.0643 (0.0005)	-0.2158 (0.0014)	-0.3147 (0.0020)	-0.0367 (0.0002)	0.1063 (0.0007)	-0.1759 (0.0051)	-0.6663 (0.0043)	0.3180 (0.0020)	0.0529 (0.0004)	0.1075 (0.0008)	0.7605 (0.0015)
7. transport	-0.3891 (0.0023)	-0.1177 (0.0007)	-0.2273 (0.0012)	0.0694 (0.0003)	-0.0435 (0.0003)	-0.1717 (0.0010)	-0.8737 (0.0009)	-0.3229 (0.0018)	-0.2961 (0.0017)	-1.5343 (0.0087)	1.2306 (0.0013)
8. education	-0.1532 (0.0012)	-0.2026 (0.0016)	-0.2066 (0.0017)	-0.0062 (0.0001)	0.0101 (0.0001)	0.1878 (0.0015)	-0.7685 (0.0062)	0.4611 (0.0116)	0.0376 (0.0003)	-0.2063 (0.0016)	0.8470 (0.0012)
9. recreation	0.0301 (0.0041)	-0.1094 (0.0004)	-0.2628 (0.0011)	-0.1258 (0.0005)	0.0536 (0.0002)	0.0041 (0.0001)	-0.3370 (0.0014)	0.0040 (0.0001)	0.0301 (0.0041)	-0.3402 (0.0014)	1.1137 (0.0004)
10. miscellaneous	-0.4728 (0.0019)	-0.0621 (0.0005)	-0.0832 (0.0005)	-0.0532 (0.0003)	-0.1199 (0.0005)	-0.0147 (0.0001)	-0.2586 (0.0012)	-0.0856 (0.0005)	-0.2071 (0.0008)	-1.3550 (0.0007)	1.6650 (0.0025)

Note: The value in parenthesis is the standard error of elasticity.

ships cannot be satisfied. Our test result supports the null hypothesis statistically, and we find that the theoretical restrictions on the economics hold in the long-run vectors. The Slutsky symmetry is not violated in the panel cointegration relationships.

Table 5 reports the long-run price and expenditure elasticities based on the cointegration estimates in the panel TECM. These are calculated by (8) and (9). The estimated expenditure elasticities are all significant at the 5% significance level. The expenditure elasticities for clothing, transport, recreation, and miscellaneous are elastic, and the expenditure effects are large. The demand for these commodities increases more than 1% for a 1% increase of real expenditure; in particular, the change for miscellaneous is large. On the other hand, expenditure elasticities for food, housing, fuel, furniture, medicine, and education are expenditure inelastic, and the effect of a 1% increase in real expenditure is small. However, the expenditure elasticities for furniture and education are relatively close to 1 and these are necessities similar to luxury goods. The estimated long-run price elasticities are all significant at the 5% level. The Marshallian own-price elasticities could be either positive or negative; however, it is usually negative. Our result shows these commodities, except for housing, clothing, and education, satisfy the negativity condition. With respect to long-run own-price elasticity, miscellaneous is price elastic, and transport is near to price elastic; in particular, the change of own-price has the largest effect on miscellaneous. If

cross-price elasticity is negative, the i th goods against the j th goods are complementary. In contrast, if cross-price elasticity is positive, the i th goods against the j th goods are substitutes. Generally, more than half the estimated elasticities are significantly negative and indicate complementarity. For example, the cross-price elasticities for miscellaneous show its complementary relationship with all cross goods. On the other hand, few substitutes are indicated. For example, both medicine and recreation are substitutes for food. That is, the rise of food prices increases demand for these items. Thus, this influence may increase the budget shares for these items to an extent. Overall, we find that many goods increase in utility by complementing other goods in the long-run.

5. Impact Measurement of Consumption Tax Hike on the Cost-of-Living Index

First, **Figure 3** shows the change in the average cost-of-living index for the base preference at prices p^0 in 1989. In addition, **Table 6** indicates the effect of the consumption tax hike in 1997 and 2014. As the cost-of-living index of the base preference in 1989 is 0, the cost-of-living indexes consistently rise from 1989 through 1998. In particular, from 1989 to 1994, the rise in the index is remarkable due to the introduction of the consumption tax and rise in prices. That is, the introduction of the consumption tax in 1989 contributes to the rise in the cost-of-living index in a time-series dimension. When the consumption tax was raised from 3% to 5% in 1997, the cost-of-living index was 0.2296 and also raised

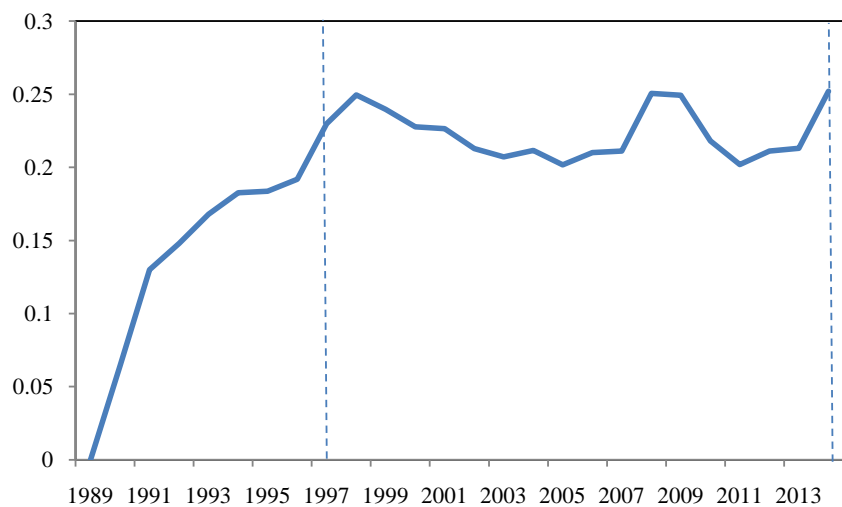


Figure 3. The average cost-of-living index from 1989-2014.

Table 6. The effect of consumption tax hike in the cost-of-index.

Year	tax in	cost-of-living index	increase rate
1997	5%	0.2296 (0.0072)	19.7%
2014	8%	0.2518 (0.0111)	18.3%

Note: The values in parentheses are standard errors.

with an increase rate of 19.7%. In other words, the consumption tax hike in 1997 was substantial to raise the cost-of-living index and influenced for household budgets. However, the rise of the cost-of-living index did not last long. After 1999, the index movement dropped by continued price deflation, but temporarily recovered in 2008. During this time, the Japanese economy had been facing a deflation problem, but the cost-of-living index itself was at a higher level than in the 1990s, and had stopped declining after that. From 2012, under the “Abenomics,” the movement of the cost-of-living index has been recovering gradually. Furthermore, the index recorded the past highest level of 0.2518 due to the consumption tax hike in 2014 to 8%, with an increase rate of 18.3%. As well as in 1997, the consumption tax hike in 2014 largely pushed up the cost-of-living index. In addition to the tax increase, it would have also been affected by the improvement of utility level with the price rise. That is, two consumption tax increases in 1997 and 2014 was effective in raising the cost-of-living index and the impact on household budgets was large. As a result, we can observe the increase in consumer surplus for 26 years because of the changing consumption expenditures and prices, added to tax increase.

Next, **Figure 4** and **Figure 5** show the difference and the increase rate in the cost-of-living index between 47 prefectural capitals in both 1997 and 2014. In general, the cost-of-living index is not very high in Tokyo, Nagoya, and Osaka, which are among Japan’s three major metropolitan cities. Rather, Tokyo is lower than other local cities. As described above, the cost-of-living index in 2014 is higher compared to 1997. In 1997, the cost-of-living index of Takamatsu is 0.3699, which is the highest value among 47 prefectural capitals. Conversely, Tokushima is 0.1347, which is the lowest value. However, looking at an increase rate in **Figure 5**, Tokushima has a high increase rate of 34.1%. In addition, local prefectural capitals such as Takamatsu, Kochi, Hiroshima, and Gifu have a high increase rate in the consumption tax hike in 1997. The cost-of-living indexes of three large metropolitan cities such as Tokyo, Nagoya, and Osaka are also in-

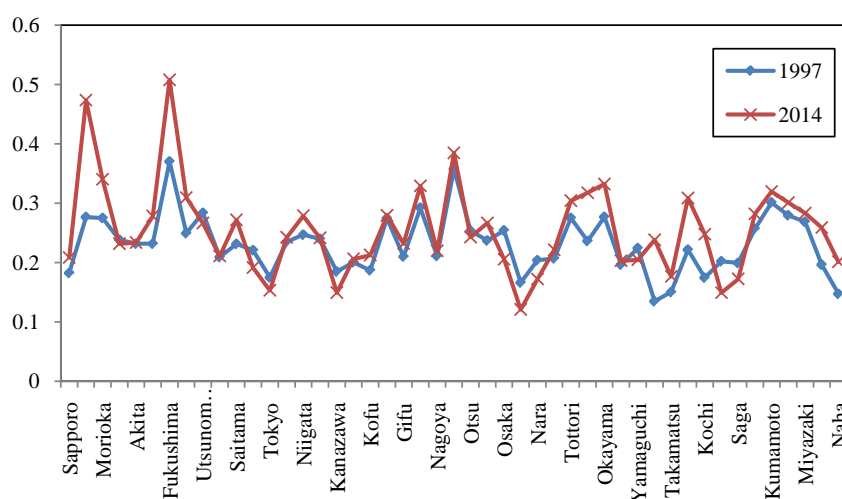


Figure 4. Cost-of-living index in 47 prefectural capitals.

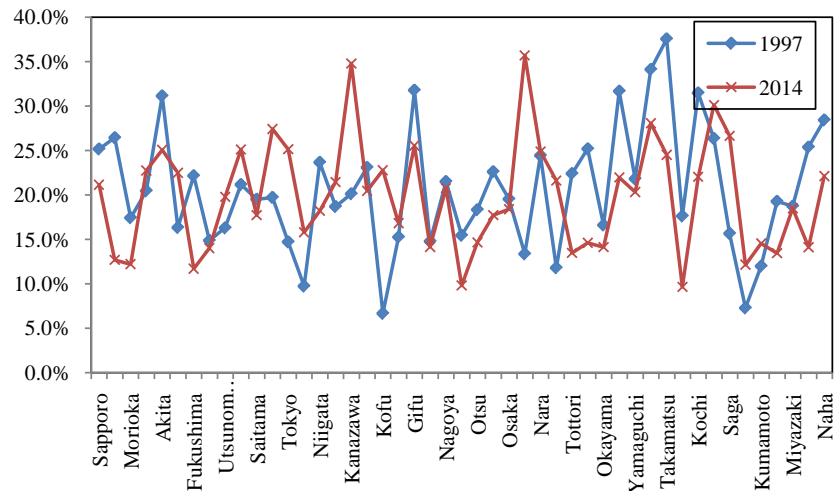


Figure 5. The increase rate of the cost-of-living index in 47 prefectural capitals.

creasing, but there is no remarkable increase as compared to other local cities. In 2014, the cost-of-living index of Fukushima is 0.5078, which is the highest value among 47 prefectural capitals. Conversely, Kobe is 0.1207, which is the lowest value. Compared to 1997, the difference in the cost-of-living index among 47 prefectural capitals is expanding in 2014. Looking at an increase rate, Kobe has the highest increase rate of 35.7% among 47 prefectural capitals. Further, Tokyo, Chiba, Kanazawa, and Fukuoka also have a high increase rate. Most of these are larger prefectural capitals located in urban areas, defined as government-ordinance-designated cities and showing the large impact of tax increase in 2014. As a difference in the effect of the tax increase between 1997 and 2014, in many cases, in 1997 the impact of tax increases was higher in local prefectural capitals than in larger prefectural capitals, but in 2014 the impact on larger prefectural capitals was higher. This difference between 1997 and 2014 is due to the differences in the reaction to price changes and differences in utility level. However, a reference about certain causes should be avoided in this study. In order to raise the effect of the tax increase on the cost-of-living index to each prefectural capital, appropriate policies of local governments need to be designed.

6. Concluding Remarks

In this study, we measured the changes in the cost-of-living index experienced due to two tax increases in Japan. The impact of the first tax increase in 1997 may be high, and therefore the rise in the cost-of-living index by the tax increase is also large, although the rise in prices was not remarkable. However, the rise of the cost-of-living index did not last long; in the last few years, it has begun to decrease. That is, the impact of the 1997 tax increase for households was lasted only for a few years. On the other hand, the impact of the second tax increase in 2014 has been also high as well as that of 1997 and the cost-of-living index has reached the highest level ever. The consumption tax increase in 2014 was effec-

tive in raising the cost-of-living index, affected by price rise and the improvement of utility level with the price rise. Therefore, we expect that this effect is not temporary unlike the case in 1997. In addition, this result indicates that Japanese households are recovering from the stagnation.

However, comparing the cost-of-living indexes among 47 prefectural capitals, we found that the three major metropolitan areas such as Tokyo, Nagoya, and Osaka have a lower cost-of-living index than other local prefectural capitals. Further, the cost-of-living indexes of these three major cities have not increased much in the tax increase of 1997, but in the case of 2014 it had a high effect. On the other hand, the local prefectural capitals obtained the effect of increasing the cost-of-living index at the 1997 tax increase, but when in 2014 it was not able to obtain a noticeable effect as before. This sluggishness in the cost-of-living index suggests that some measures for prices and household budgets need to be implemented in each prefectural capital. It will bring about improvement of the cost-of-living index and have further effect on future tax increases.

The long-term stagnation in prices was problematic for the Japanese economy, but we found that the cost-of-living index increased over the 26 years studied here. Indeed, the tax increase has improved the cost-of-living index from the previous stagnant state, and it seems that the timing of two tax increases in both 1997 and 2014 was appropriate. In particular, the effect on the tax increase in 2014 is expected to the future recovery for household budgets. As a result, we had different effects between 1997 and 2014 in this study. It would be desirable to conduct a detailed analysis on the cause of this difference in the future. By clarifying this cause, it seems that it will become possible to increase the effect on future consumption tax increase. In addition, to measure the effects of the consumption tax increase more accurately, it is desirable to be able to remove the direct influence of consumption tax on prices. This task is difficult at the present study and should be required for future research.

References

- [1] Deaton, A. and Muellbauer, J. (1980) An Almost Ideal Demand System. *American Economic Review*, **70**, 312-326.
- [2] Ogura, M. (2015) The Estimation of a Panel Cointegrated Demand System in Japan: Results from a Panel Vector Error Correction Model. *International Journal of Intelligent Technologies and Applied Statistics*, **8**, 117-132.
- [3] Stone, R. (1954) *The Measurement of Consumers' Expenditure and Behavior in the United Kingdom, 1920-1938*. Vol. 1, Cambridge University Press, Cambridge.
- [4] Attfield, C.L.F. (1997) Estimating a Cointegrating Demand System. *European Economic Review*, **41**, 61-73. [https://doi.org/10.1016/0014-2921\(95\)00062-3](https://doi.org/10.1016/0014-2921(95)00062-3)
- [5] Ogura, M. (2011) Testing the Structural Break in Japanese Expenditure Patterns after the Bubble Era. *Structural Change and Economic Dynamics*, **22**, 277-286. <https://doi.org/10.1016/j.strueco.2011.03.003>
- [6] Green, R.D. and Alston, J.M. (1990) Elasticities in AIDS Models. *American Journal of Agricultural Economics*, **72**, 442-445. <https://doi.org/10.2307/1242346>
- [7] Buse, A. (1994) Evaluating the Linearized Almost Ideal Demand System. *American*

- Journal of Agricultural Economics*, **76**, 781-793. <https://doi.org/10.2307/1243739>
- [8] Fry, V. and Parshardes, P. (1989) Constructing the True Cost of Living Index from the Engel Curves of the PIGLOG Model. *Journal of Applied Econometrics*, **4**, 41-56. <https://doi.org/10.1002/jae.3950040104>
- [9] Lewbel, A. (1989) Household Equivalence Scales and Welfare Comparisons. *Journal of Public Economics*, **39**, 377-391. [https://doi.org/10.1016/0047-2727\(89\)90035-2](https://doi.org/10.1016/0047-2727(89)90035-2)
- [10] Pollak, R.A. (1989) The Theory of the Cost-of-Living Index. Oxford University Press, New York, Oxford.
- [11] Duffy, M. (2002) On the Estimation of an Advertising-Augmented, Cointegrating Demand System. *Economic Modeling*, **20**, 181-206. [https://doi.org/10.1016/S0264-9993\(01\)00091-8](https://doi.org/10.1016/S0264-9993(01)00091-8)
- [12] Lewbel, A. and Ng, S. (2005) Demand Systems with Nonstationary Prices. *The Review of Economics and Statistics*, **87**, 479-494. <https://doi.org/10.1162/0034653054638283>
- [13] Maddala, G.S. and Wu, S. (1999) A Comparative Study of Unit Root Tests with Panel Data and New Simple Test. *Oxford Bulletin of Economics and Statistics*, **61**, 631-652. <https://doi.org/10.1111/1468-0084.61.s1.13>
- [14] Kao, C. (1999) Spurious Regression and Residual-Based Tests for Cointegration in Panel Data. *Journal of Econometrics*, **90**, 1-44. [https://doi.org/10.1016/S0304-4076\(98\)00023-2](https://doi.org/10.1016/S0304-4076(98)00023-2)
- [15] Phillips, P.C.B. (1991) Optimal Inference in Cointegrated Systems. *Econometrica*, **59**, 283-306. <https://doi.org/10.2307/2938258>
- [16] Phillips, P.C.B. (1994) Some Exact Distribution Theory for Maximum Likelihood Estimators of Cointegrating Coefficients in Error-Correction Models. *Econometrica*, **62**, 73-93. <https://doi.org/10.2307/2951476>
- [17] Newey, W.K. and West, K.D. (1987) A Simple, Positive Semi-definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix. *Econometrica*, **55**, 703-708. <https://doi.org/10.2307/1913610>



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